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Evidence on Real Exchange Rate – Inflation Causality: An Application of Toda-Yamamoto Dynamic Granger Causality Test

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ABSTRACT

The paper provides further evidence of the real exchange rate and inflation causal relationship using Toda-Yamamoto (1995) augmented Granger causality test in Malaysia, Nigeria, Philippines and South Africa. The critical values used in this study are simulated based on the leverage bootstrap. The results are compared between the Granger asymptotic chi-square distribution, the modified WALD test statistics and the leverage bootstrapped distribution critical values. Conflicting findings are obtained which prove the existence of size distortion and nuisance parameter estimates when the former method is applied. The result based on the Toda-Yamamoto and leverage bootstrapped critical values reveal that policy intervention on inflation can stabilize real exchange rate in Malaysia and Nigeria but not vice versa. Moreover, bidirectional causation exists in Philippines and South Africa meaning that any policy intervention formulated on one variable can stabilize the other. The policy implication of this finding is that the policy makers can manipulate the rate of inflation to stabilize real exchange rate fluctuations in all countries under study, but can only regulate inflation through exchange rate in the case of Philippines and South Africa.

Keywords: Leverage bootstrap, Toda-Yamamoto causality, Real exchange rate, Inflation, Structural break.

1. Introduction

The nature of causality between real exchange rate and inflation has been the subject of concern in international economics and macroeconomics. Conflicting findings are reported by researchers that suffer from methodological problems and hamper appropriate policy formulation. The question is whether inflation influences real exchange rate or the other way round, that is, whether a variable can be influenced by the past observations of another variable, refer to as Granger definition of causality (Granger, 1969). This causal relationship is re-investigated using a modified leverage bootstrap distribution and Toda-Yamamoto causality. The methods work better under small sample size, violation of the normality assumption and in the presence of autoregressive conditional heteroscedasticity (ARCH) effect (Hatemi-J & Irandoust, 2006). Previous studies were confronted with invalid inferences, as a result, of testing Granger causality based on Johansen Juselius co-integration, vector error correction and ordinary unrestricted vector autoregressive model with inappropriate lag length. This leads to nuisance parameter estimates (Guru Gharana, 2012).

The traditional Granger causality was developed on asymptotic distribution theory which is seen to results to a spurious conclusion when variables are integrated of order one I (1) (Granger & Newbold, 1974). Most at times, the methods of modelling causality in the traditional sense were found to have overlooked some elements of the forecast that results to invalid results of non-granger causality. This implies that in some cases, causality exist but spuriously not found due to the inefficiency of techniques employed (Granger, 1988). Another limitation of Granger causality is that the null hypothesis at level estimation suffer from non-standard asymptotic distribution, whereas, the integrated Granger causality suffer from independence of nuisance parameter estimates (Sim, Stock & Watson, 1990 and Toda & Philips, 1993). Furthermore, the hypothesis of vector error correction model of Granger non-causality applies nonlinear parameter restriction matrices. The Wald and likelihood ratio test statistics for Granger test are associated with rank deficiency which leads to size distortion under null hypothesis (Toda & Philips, 1993). The Granger method also mandates testing for stationarity and co-integration properties which are less relevant if the goal for the test is to focus on the causal relationship between variables (Hacker & Hatemi-J, 2006 and Lee, Lin & Wu, 2002).

However, in previous studies, the unit root properties have mostly been checked using Augmented Dickey-Fuller (ADF) (1979) test based on the assumption that present shocks are only temporal and do not have long run effect on the series. This is proved to be deficient and results to bias and less power to reject the null hypothesis if structural breaks exist (Perron, 1989 and Zivot & Andrews, 1992). This study employs Lee and Strazicich (2003) minimum lagrange multiplier with a structural break to test the integration properties for the purpose of determining the maximum order of integration among the series. The test solves the problem of weak assumption of absence of a break in the null hypothesis associated with endogenous break determined test such as Zivot and Andrews (1992), Lumisdaine and Pappel (1997) and Perron (1997) as well as other similar tests. The assumption results to significant rejection of the unit root with break (Lee & Strazicich, 2001 and Nunes, Newbold & Kuan, 1997).

In the previous endogenous tests, rejection of the null hypothesis does not indicate trend stationarity, rather a unit root rejection with break, whereas, in the Lagrange Multiplier (LM) test rejection of the null hypothesis satisfy stationarity in trend. The test is, break point nuisance invariant under both null and alternative hypothesis suffering neither size nor location distortion. The test considers two different breaks under the unit root null without relying on the nuisance parameter. This makes the test free from spurious rejection and unaffected by size and incorrect estimation irrespective of whether structural break is present or not (Lee & Strazicich, 2004). The test is also efficient in addressing the problems of divergence, as a result of an increase in the magnitude of the break due to independence of the test on break location.

It is also argued that causality should be tested within a jurisdiction of an acceptable theory. This assists in determining appropriate variables to be included in the model which will capture the required information on causality (Zellner, 1978 in Granger, 1980). In addressing the prevailing inconsistencies in the causal direction between real exchange rate and inflation, a conditional causality is tested based on monetary theory of exchange rate determination.

The present study employs the potent Toda & Yamamoto (1995) modified Wald test statistics based on augmented Vector Autoregressive (VAR) framework that applies the asymptotic $\chi2$ distribution to ease the difficulties encountered in Granger test of causality. Their approach is applicable irrespective of whether series are stationary at level I (0), integrated of the same order I (1), arbitrarily integrated or co-integrated of arbitrary order. The procedure also tests for coefficient linearity and non-linearity restriction through Wald criterion application on estimated level VAR (Hacker & Hatemi-J, 2006). Moreover, the lag length selection method used in the level VAR is appropriate for all VAR estimates including integrated and co-integrated vector autoregressive (VAR) process provided that the lag length is equal or greater than the order of the integration (Toda & Yamamoto, 1995).

However, Sim, Stock & Watson (1990) against the application of asymptotic distribution theory on a VAR model to test for causality among level and integrated variables even if they are cointegrated. It leads to size distortion under small sample size as ours. Nevertheless, bootstrapped distribution is considered more reliable than asymptotic distribution in a finite sample study in order to avoid size distortion and spurious inferences (Hacker & Hatemi-J, 2006). This is first suggested by Efron (1979) who furnishes more reliable critical values especially for small sample analysis.

Therefore, this study test for the order of integration using Lee and Strazicich (2003) to account for structural breaks. The paper combines the robust Toda-Yamamoto (1995) causality and Hacker and Hatemi-J (2006) leverage bootstrap distribution theories. This will address the inadequacies of the past studies that employed Granger (1969) causality approach, keeping in mind the end goal to better the soundness and consistency of the inferences and test whether the result will be different. The combination of these rigorous approaches has not been used in the context of causation between inflation and real exchange rates.

The other sections of the paper are planned under four headings. Section two reviews previous literature on the causal relationship between exchange rate and inflation. Section three deals with the theoretical framework. Section four describes the data. Section five deals with methodology and empirical result and section six present conclusion and policy implication.

2. Literature Review

In the literature, empirical findings on the causal relationship between exchange rate and inflation have been inconsistent using same methodology (Granger causality), same data generating process as well as different methodologies and different sample sizes. In some studies, bidirectional causal relationship between the variables was discovered. These include the studies of Arabi (2012); Arize and Malindretes (1997) & Madesha, Chidoko and Zivanomoyo (2013).

Arabi (2012) & Madesha, Chidoko and Zivanomoyo (2013) estimate exchange rate volatility under the presence of GARCH effect. The results reveal a simultaneous long run feedback possibility between exchange rate fluctuations and its determinants such as inflation. In another word, inflation and exchange rate volatility granger cause each other. Nonetheless, the presence of ARCH effect makes Granger test inefficient (Hacker & Hatemi-J, 2006). Arize and Malindretes (1997) study exchange rate volatility as a factor that causes variability in inflation in 41 countries. The result confirms the existence of bidirectional or simultaneity between exchange rate variability and inflation under the flexible exchange rate era.

On the contrary, Achani, Jayanthy, Fauzi and Abdullah (2010); Imimola and Enoma (2011) & Omotor (2008), establish a unidirectional relationship from exchange rate to inflation. Achani *et al*., (2010) examine the effect of inflation and exchange rates in Asean + 3, the EU and the North America. They discover one-way causality in Asia with a strong correlation between inflation and exchange rate in most of the Asian countries except Malaysia. They further find that the sensitivity of inflation to fluctuations in the exchange rate to be higher in Asia than E U and North America. Likewise, Imimola and Enoma (2011) & Omotor (2008) explore the effect of exchange rate depreciation on inflation in Nigeria. They find inflation to be largely determined by exchange rate depreciation, money supply and output.

However, some studies particularly in Nigeria, show no causality between inflation and exchange rate (Cairns, Ho & McCaulay, 2007; Chen & Wu 2001; Emmanuel, 2013; Kamas, 1995; Nnamdi & Ifionu, 2013; Parvar, Mohammed & Hassan, 2011).

Emmanuel (2013) examines the effect of foreign exchange reserve on exchange rate and inflation in Nigeria. The study reveals a significant causal relationship between foreign exchange reserves and exchange rate that influences the volatility or otherwise of Naira/USD exchange rate without any causation from neither inflation nor exchange rate. Nnamdi and Ifionu (2013) examine volatility in the exchange rate. The findings of their study indicate that exchange rate volatility is significantly influenced by the uncertainty in the lagged exchange rate independent of inflation and other variables. Chen and Wu (2001) analyze real exchange rate fluctuations sources in four pacific basin countries. Their study shows that a real shock (technology, resource endowment, preferences) impact positively on real exchange rate fluctuations, unlike in other studies where it is assume to be influenced by inflation.

However, these researches suffer from invalid inferences as a result, of testing Granger causality based on Johansen Jesulius co-integration, vector error correction, ordinary least squares (OLS) and vector autoregressive model with inappropriate lag length. This leads to non-standard asymptotic distribution, nuisance parameter estimates and rank deficiency which results to size distortion under null hypothesis (Guru Gharana, 2012; Toda & Philips, 1993). In another development, Sim, Stock and Watson, (1990) against the application of asymptotic distribution theory on a VAR model to test for causality among level and integrated variables even if the variables are cointegrated.

3. Theoretical Framework

There exist various theories that clarify the causal linkage between exchange rate disequilibrium and inflation. In this paper, the theoretical framework underpinning the study is the monetary theory of exchange rate determination divided into sticky-price monetary model developed by Dornbusch (1976) and flexible-price version introduced by Frankel (1976) and Mussa (1976).

The theory of exchange rate started with Purchasing Power Parity (PPP) or what is called “inflation theory of exchange rate” without which exchange rate disequilibrium cannot be determined (Cassel, 1918). It is denoted as partial equilibrium theory because of its inability to explain the phenomenon of money market and balances of foreign payment in the determination of exchange rate (Kanamori & Zhao, 2006). The monetary approach to exchange rate determination explains the significance of money and other variables (assets) in defining the factors responsible for determining exchange rate under flexible regime and balance of payment under pegged regime (Frenkel, 1976 in Frenkel & Johnson, 2013). The equilibrium exchange rate is obtained when demand for and supply of money are held willingly. Although, some researchers recognized that exchange rate and inflation are simultaneously determined, Cassel (1921) as emphasized in Frenkel (1976) and Whitney (1922) that causation exist between the variables. Cassel (1921) argues that there is a flow of causality from price (inflation) to exchange rate, whereas, Einzig (1935) claims that the causation runs from exchange rate to inflation.

The monetary theory depicts the relationship between exchange rate and inflation, interest rate, money supply and real income. The theory postulates that increase in money supply increases domestic inflation rate and increase in domestic inflation lead to low demand for local currency that causes high domestic exchange rate. This means that any change in the supply of money has a proportionate effect on exchange rate stability. It is expected that increase in domestic money supply causes a corresponding increase in exchange rate. The monetary model provides that interest rate is negatively related to exchange rate under the sticky-price model. This is because raises in the domestic rate of interest influences incipient capital into the economy that leads to high demand of local currency. Whereas, under the flexible price model, increase in interest rate cuts the demand for local currency which causes currency depreciation in the domestic economy (Frankel, 1982).

4. Data

For the purpose of this study, annual data from 1980 to 2012 was collected on real effective exchange rate, inflation, interest rate and money supply from World Development Indicators (WDI) and Financial Statistics of International Monetary Fund (IMF) for Malaysia, Nigeria, Philippines and South Africa.

The researchers are mainly concern about the direction of causality between inflation and exchange rate. That is the amount to which one variable scientifically influence changes in another variable by observing the connectedness of previous values of a particular variable on another variable. The other variables aside exchange rate and inflation are regarded as controlled variables in the model. Control variables play a significant role of ensuring presence or absence of causation between variables under investigation where causality might have been absent or present respectively. If control variables are found to have a significant influence in causing a relationship, it will be very vital in decision making process regarding how to control the effect (Granger, 1980).

5. Methodology

This paper applies leverage bootstrap distribution developed by Hacker and Hatemi-J (2006) on Toda and Yamamoto (1995) causality approach. The integration order of the variables was tested using Lee and Strazicich (2003) to account for structural breaks and address the shortcomings of the previous studies that employed various unit root test such as ADF and PP among others. The paper also aims at solving the highlighted problems associated with Granger (1969) causality approach and improve the soundness of the inferences to be drawn.

5.1. Lee and Strazicich Unit Root Test

Although, the unit root properties is not a major concern in this methodology because of the applicability of the approach irrespective of whether series are stationary at level I(0), integrated of the same order I(1), arbitrarily integrated or co-integrated of arbitrary order (Hacker & Hatemi-J, 2006 and Lee, Lin & Wu, 2002). This paper adopts the robust Lee and Strazicich (2003) minimum lagrange multiplier with structural break to test the integration properties for the purpose of determining the maximum order of integration among the series. The test solves the problem of weak assumption of absence of a break in the null hypothesis associated with endogenous break determined test. The test is break point nuisance invariant under both null and alternative hypothesis unaffected by neither size nor location distortion. This makes the test free from spurious rejection and unaffected by size and incorrect estimation irrespective of whether structural break is present or not (Lee & Strazicich, 2004). The unit root is tested using the following equations:

$$y_t = \delta Z_t + X_{t,} \qquad X_t = \beta X_{t-1} + \varepsilon_t \qquad [1]$$

Where Z_t is a vector that contains exogenous variables. The null hypothesis is given as $\beta = 1$, ε_t is a white noise process that is, $\varepsilon_t \sim iidN(0,\sigma^2)$ which can be relaxed to ensure the absence of autocorrelation. The test cogitates two different structural changes. The first model permits dual changes in intercept defined as Z_t = [1, t, D_{1t}, D_{2t}]' in which D_{jt} = 1 for time $t \geq T_{Bj} + 1, j = 1,2$ and zero otherwise. The second model takes account of two shifts in intercept and trend. This is defined as Z_t = [1, t, D_{1t}, D_{2t} $DT^*_{1t} DT_{2t}$]' $DT^*_{jt} = t - T_{Bj}$ *for* t ≥ $T_{Bj} + 1, j = 1,2$ and zero otherwise. T_{Bj} indicates the break time period. Recall that the break period in this test is embedded under both null and alternative hypothesis as $\beta = 1$ and $\beta < 1$ respectively as shown below:

Null hypothesis $\qquad y_t = \mu_0 + d_1 B_{1t} + d_2 B_{2t} + y_{t-1} + v_{1t} \qquad [2]$

Alternative hypothesis $\qquad y_t = \mu_1 + \gamma_t + d_1 D_{1t} + d_2 D_{2t} + y_{t-1} + v_{2t} \qquad [3]$

v_{1t} and v_{2t} are white noise errors, B_{jt} = 1 for t = $T_{Bj} + 1, j = 1,2$ and zero otherwise, d = (d_1 d_2). To ensure time invariant test to the break size, that is why dummy variable B_{jt} is included in equation two (Perron, 1989). Similar models with additional D_{jt} and DT_{jt} terms to equation two and three respectively should be repeated to depict the models that allow for two shifts in both intercept and trend.

The following lagrange multiplier regression equation will be estimated to derive the two breaks unit root $t-$statistics.

$$\Delta y_t = \delta \Delta Z_t + \phi \tilde{X}_{t-1} + \mu_t \qquad [4]$$

Where $\tilde{X}_t = y_t - \tilde{\psi} - Z_t \tilde{\delta}$, $t = 2,\ldots,T; \tilde{\delta}$ is a coefficient of Δy_t on ΔZ_t; $\tilde{\Psi}_x$ is $y_t - Z_t \tilde{\delta}$. The null hypothesis is defined as $\phi = 0$ and the langrange multiplier $t-$statistics is given as:

$$\tilde{\rho} = T.\tilde{\phi} \qquad [5]$$

$\tilde{\tau} =$ $t-$statistics for testing the unit root of LM test $\phi = 0$ [6]

The minimum LM two breaks unit root determines the break period (TB_j) as written below:

$$LM_p = \inf_{\lambda} \tilde{\rho}(\lambda) \qquad [7]$$

$$LM_\tau = \inf_{\lambda} \tilde{\tau}(\lambda) \qquad [8]$$

The break periods are indicated by the minimum t-statistics.

5.2. Toda - Yamamoto Causality

For the test of a causal relationship between the variables, the study uses the powerful Toda and Yamamoto (1995) modified Wald test statistics based on augmented Vector Autoregressive $VAR(p+d)$ framework.

$$y_t = \alpha_0 + \alpha_1 t + \ldots + \alpha_q t^q + \eta_t \qquad [9]$$

η_t is k^{th} order of vector autoregressive (VAR) process which is further defined as $I(d)$ and possibly $CI(d,b)$.

$$\eta_t = j_1 \eta_{t-1} + \ldots + j_K \; \eta_{t-k} + \mu_t \qquad [10]$$

k in equation 10 represent a known lag length. Therefore,

$$y_t = \gamma_0 + \gamma_1 t + \ldots + \gamma_q t^q + j_1 y_{t-1} + \ldots + j_K y_{t-k} + \varepsilon_t \qquad [11]$$

Where $y_{t,}$ γ and ε_t are vectors of $n-$dimension with error vector that is $\sim iid\, N(0,\sigma^2)$ and covariance matrix that satisfy the condition that $\Sigma_\varepsilon > 0$ and that $E\,|\,\varepsilon_{it}\,|^{2+\lambda}$ less than infinity where δ is greater than zero.

The restricted hypothesis is formulated as; $Ho = f(J) = 0$

The estimated form of the equation on which the hypothesis will be tested is given in equation 12 below:

$$y_t = \hat{\gamma}_0 + \hat{\gamma}_1 t + \ldots + \hat{\gamma}_q t^q + \hat{j}_1 y_{t-1} + \ldots + \hat{j}_K y_{t-k} + \hat{j}_p y_{t-p} + \hat{\varepsilon}_t \quad [12]$$

From the above equation, y_t is a vector of real exchange rate, inflation, interest rate and money supply and d lag more than the true lag length k is included. The parameters of the added lags are left unrestricted in the null hypothesis test to ensure the effectiveness of the asymptotic chi-square values when normality assumption is fulfilled in the VAR model (Toda & Yamamoto, 1995). Hence, equation 12 is deduced as:

$$y_t = \hat{\Gamma}\tau_t + \hat{\phi} x_t + \hat{\psi} z_t + \hat{\varepsilon}_t \quad [13]$$

Where $\hat{\Gamma} = (\gamma_0, \ldots\ldots, \gamma_q)$, $\hat{\phi} = (\hat{j}_{1,} \ldots\ldots, \hat{j}_K)$, $\hat{\psi}$ = ($\hat{j}_{K+1,} \ldots\ldots, \hat{j}_p$), τ_t = (1, t, …, t^q), x_t = ($y'_t, \ldots, y'_{t-k}$), z_t = ($y'_{t-k-1,} \ldots, y'_{t-d}$)$'$ or can be represented in a regular matrix form as in equation 14 below:

$$Y' = \hat{\Gamma}\tau' + \phi X' + \psi Z' + \sigma' \quad [14]$$

Unrestricted regression is estimated from equation (14) in order to get a vector of estimated residuals from which the variance-covariance matrix of the residuals is calculated. From equation 14 above, the null hypothesis can be tested using the constructed Wald statistics W below to test the Granger non-causality between the variables:

$$MWALD = f\left(\hat{\phi}\right)' [F\left(\hat{\phi}\right)' \{\hat{\Sigma}_\tau \otimes (X'QX)^{-1}\} F\left(\hat{\phi}\right)']^{-1} f\left(\hat{\phi}\right) \quad [15]$$

Here F is an indicator matrix $p \times n\left(1 + n\left(p+d\right)\right)$ indicating zero value parameters, the symbol $\otimes$ denotes a matrix multiplication of an element by all elements. Chi-square asymptotic critical values are employed for general restrictions on parameter matrices ($\hat{j}_{1,} \ldots\ldots, \hat{j}_K$) to test the null of no Granger causality with restrictions equal to the degrees of freedom. This will be done by estimating an over-fitted model after determining the maximum order of integration d_{max}. The lag order to be included will be $p = k + d_{max}$. Assuming y_t is found to be integrated of order two with a linear trend. Then, the following equation will be estimated:

$$y_t = \hat{\gamma}_0 + \hat{\gamma}_1 t + \hat{j}_1 y_{t-1} + \ldots + \hat{j}_{K+1} y_{t-k-1} + \hat{j}_{k+2} y_{t-k-2} + \hat{\varepsilon}_t \quad [16]$$

Here under the restricted hypothesis, there exist asymptotic distribution using the normal degree of freedom without regard to stationarity, integration and co-integration order of the variables. Moreover, the lag length selection method used in the level VAR is appropriate for all VAR estimates including integrated and co-integrated VAR process provided that the lag length is greater than or equals to the order of the integration ($k \geq d$). The correct lag length is chosen through testing the significance of $J_{k+1,}\ J_{k+2,} \ldots J_p$ in equation (16) for $p > k$ condition (Toda & Yamamoto, 1995) and minimizing the Hatemi-J (2003) information criterion described underneath.

$$HJC = ln\left(\left|\bigcap_z\right|\right) + z \times v^2 \left(\frac{\ln N + 2\ln\left(lnN\right)}{2N}\right) \quad z = 0, \ldots, p. \quad [17]$$

Where HJC is the Hatemi-J information criterion, ln is the natural logarithm, $\left|\bigcap_z\right|$ represent the lag order z determinant of the estimated white noise variance-covariance matrix in the VAR framework, v and N denote the number of variables and observations used in the VAR model respectively. Furthermore, equation (17) has been tested to work better especially if integration exist among the variables (Hatemi-J, 2003).

However, when normality assumption is not met, and the effect of autoregressive conditional heteroscedasticity exist, the usual asymptotic distribution theory does not work well (Hatemi-J & Irandoust, 2006). Therefore, more reliable leverage distribution theory will be employed in this kind of finite sample to avoid size distortion and spurious inferences.

5.3. Leverage Bootstrapping

The leverage bootstrap critical values are generated with GAUSS using the program procedure developed by Hacker and Hatemi-J (2010). Therefore, following Hacker and Hatemi-J (2006) the leverage bootstrap simulation is conducted in the following way. Equation (14) is estimated with restriction of no Granger causality, and simulated data is generated for every bootstrap simulation

$$y_t^* = \hat{\Gamma}\tau_t + \hat{\phi}x_t + \hat{\psi}z_t + \varepsilon_t^* \qquad [18]$$

Where the bootstrap residuals ε_t^* are estimated on the basis of N random draws with replacement from equation (18) modified residuals with identical probability of $1/N$ in each case. To set the bootstrapped residuals mean value equals zero, the mean value of the modified residuals is deducted from each modified set of residuals. Leverage adjustment is employed to adjust the raw residuals of the regression to arrive at constant variance. Hacker and Hatemi-J (2003) define leverage adjustment for multivariate functions such as y_{it} as follows:

$$\varepsilon_{it}^m = \frac{\varepsilon_{it}}{\sqrt{1-h_{it}}}, \qquad [19]$$

Here ε_{it} is the ordinary residuals obtained from the y_{it} $(i=1,2,3,\ldots 4)$ regression, h_{it} is the h_i t^{th} element defined as $h_i = \text{diag}\left(Y_1\,(Y_1^{'}Y_1)^{-1}\,Y_1^{'}\right)$ and h_j = diag $\left(Y\,(Y_1^{'}Y_1)^{-1}\,Y'\right)$ for $i=1,2,3,\ldots 4$ and $j=i-1$. Where Y_1 is a regression matrix of independent variables that determine y_{1t} with no Granger causality restriction and Y is a set of the regression matrix of regressors that explain y_{jt} including the lags of all variables in the estimation. This scenario prevails when testing the null hypothesis of no Granger causality from y_{jt} to y_{1t} and vice versa.

In this study, the critical values are generated based on the underlying empirical data through bootstrap simulation. The iteration is conducted 1000 times and *MWALD* $t-$ statistics are estimated after every iteration to determine the upper $(\alpha)^{th}$ quantile of the bootstrapped distribution of the *MWALD* $t-$ statistics in order to generate 1%, 5% and 10% bootstrapped critical values. Finally, the raw data rather than the bootstrapped one is utilized to calculate the *MWALD* statistics. The hypothesis of no Granger causality is rejected if the *MWALD* statistics calculated using the original data is greater than the bootstrapped critical values (C_α^*).

5.4. Empirical Result

The integration properties of the series is investigated using LS test (Lee & Strazicich, 2003) to determine the data generating process of the variables. The result is presented in table one below:

Table-1. Lee and Strazicich Two-Break Minimum Lagrange Multiplier (LM) Unit Root Test

	Model A					Model C				
Variables	k	$\hat{T}_B$	$\hat{t}_{\gamma j}$	***Test Statistic***	***Critical Value Break Points*** λ	k	$\hat{T}_B$	$\hat{t}_{\gamma j}$	***Test Statistic***	***Critical Value Break Points*** λ
MALAYSIA *RER*	1	1987	-1.203	-2.731	-.04	1	1987	-3.847*	-5.974*	-.12
		2009	1.268		.04		1997	1.340		.04
INF	1	1984	-5.345*	-3.833***	-.17	1	1986	4.116*	-4.095	.13
		1987	3.487*		.11		1993	-4.148*		-.13
INT	1	1994	1.373	-3.909**	.04	1	1985	-2.942**	-4.683	-.09
		1998	-3.247*		-.10		1998	1.251		.04
MSS	1	1995	1.770***	-3.315	.06	1	1991	2.932**	-4.422	.09
		1998	3.456*		.11		1998	-2.553**		-.08
										Continue

NIGERIA *RER*	1	1984	-3.085*	-3.075	-.10	1	1988	-2.660**	-4.775	-.09
		1994	-1.858***		-.06		1998	1.506		.05
INF	1	1986	-.430	-5.515*	-.01	1	1992	1.596	-6.446*	.05
		1996	-1.774***		-.06		1997	-.631		-.02
INT	1	1986	1.846***	-2.567	.06	1	1993	-3.038*	-5.612**	-.10
		1989	1.802***		.06		2000	.671		.02
MSS	1	1988	-1.419	-2.583	-.05	1	1984	.219	-5.424***	.01
		2009	1.244		.06		1991	4.307*		.14
PHILIPPINES *RER*	1	1985	-3.951*	-2.642	.13	1	1990	2.785**	-3.417	.09
		1997	-2.075**		.07		1999	-2.562***		-.08
INF	1	1984	1.280	-6.519*	.04	1	1984	-3.732*	-7.760*	-.12
		1987	-.756		-.02		1988	6.256*		.20
INT	1	1994	-1.707***	-3.905**	-.06	1	1994	-.266	-5.127***	-.01
		2001	-3.117*		-.10		2007	-2.965**		-.10
MSS	1	1988	1.874***	-2.474	.06	1	1987	4.717*	-5.023***	.15
		1992	.692		.02		2000	-2.814*		.09
	Model A					Model C				
Variables	k	$\hat{T}_B$	$\hat{t}_{\gamma j}$	*Test Statistic*	*Critical Value Break Points* λ	k	$\hat{T}_B$	$\hat{t}_{\gamma j}$	*Test Statistic*	*Critical Value Break Points* λ
S/AFRICA *RER*	1	1987	-1.817***	-4.050**	-.06	1	1987	1.688***	-4.986	.05
		1997	-1.175		-.04		1999	-2.455**		-.08
INF	1	1984	1.583	-4.904*	.05	1	1989	-.573	-6.771*	-.02
		2005	-1.899***		-.06		2005	3.296*		.11
INT	1	1998	-1.584	-3.651***	-.05	1	1988	-5.164*	-6.004*	-.17
		2003	-2.440**		-.08		1999	1.993***		.06
MSS	1	1990	1.988***	-1.901	.06	1	1995	-.370	-5.033***	-.01
		2000	-4.850*		-.16		2001	-1.344		-.04
Critical values		1%	5%	10%						
Model A		-4.545	-3.842	-3.504						
Model C		-5.823	-5.286	-4.989						

Note: k is the optimal number of lagged first-difference terms included in the unit root test to correct for serial correlation. $\hat{T}_B$ denotes the estimated break points. $\hat{t}_{\gamma j}$ is the *t* value of DT_{jt}, for j=1,2. See J. Lee and Strazicich (2003) Table 2 for critical values. a, b and c indicates significance of the LM test statistics at 99%, 95% and 90% critical level, respectively. While *, ** and *** indicates the two-tailed significance level of the break date at 99%, 95% and 90% respectively.

The above table present the unit root test conducted on the variables. The significance of this test is to know the highest order of integration of the variables. The test rejects the null hypothesis that series are *I*(2) at all levels of significance. Therefore, the maximum order of integration of all the variables for all countries is found to be *I*(1) order. This signifies that the lag augmentation (d_{max}) in estimating Toda-Yamamoto (1995) vector autoregressive model for all the countries is determined as one.

Table-2. Test for ARCH effect and normality in the VAR model

Country	ARCH effect	Normality
Malaysia	0.1891	0.0059*
Nigeria	0.0289**	0.0079*
Philippines	0.0517**	0.0223**
South Africa	0.1058	0.0000*

* and ** represent rejection of the null hypothesis at 1% and 5% significant level respectively.
Source: Authors computation using EVIEWS 8

The table 2 above depict the autoregressive conditional heteroscedasticity effect and normality in the VAR model. The null hypothesis of normality in the VAR model is rejected for all countries under study. Furthermore, the null hypothesis of non-existence of multivariate autoregressive conditional heteroscedasticity (ARCH) effect is also rejected for Nigeria and Philippines whereas, the hypothesis cannot be rejected for Malaysia and South Africa. However, the failure to fulfill the normality assumption and the existence of autoregressive conditional heteroscedasticity effect, render the usual asymptotic distribution theory to be less relevant (Hatemi-J & Irandoust, 2006). Therefore, more reliable leverage distribution theory which perform better in the presence of non-normality and ARCH effect is employed in this study.

Table-3. The result of Granger, Toda-Yamamoto causality and Bootstrap simulation

				Leverage Bootstrap		
The null hypothesis	**Non-Granger causality**	**MWALD test statistics**	**MWALD *P*. values**	**1% CV**	**5% CV**	**10% CV**
Malaysia						
INF ≠> RER	12.727(0.000)	7.694**	0.021	8.215	4.323	3.164
RER ≠> INF	0.696(0.404)	2.363	0.307	7.026	4.193	2.936
Nigeria						
INF ≠> RER	4.251(0.039)	7.868**	0.049	10.133	4.962	3.259
RER ≠> INF	4.149(0.042)	1.300	0.729	9.951	4.189	2.833
Philippines						
INF ≠> RER	5.043(0.025)	6.908**	0.032	8.425	4.283	2.809
RER ≠> INF	0.316(0.574)	9.752*	0.008	7.893	4.319	3.014
South Africa						
INF ≠> RER	5.525(0.019)	7.252**	0.027	8.621	4.320	3.002
RER ≠> INF	1.164(0.281)	6.999**	0.030	7.959	4.798	3.001

1. *, ** & *** represent rejection of null hypothesis at 1%, 5% and 10% significant level respectively, with reference to bootstrap simulated critical values.
2. The symbol ≠> represent Granger non-causality.
3. The estimated order of the VAR $(\boldsymbol{P}+\boldsymbol{d}_{max})$ model is determined to be two for all countries except for Nigeria which is determined as three. This is made up of the VAR order $\boldsymbol{P}$ and a constant one lag augmentation, because the maximum order of integration does not exceed one for all series.
4. The figures enclose in parenthesis under column two represent the p values of Granger non-causality.

Source: Authors computations using RATS and GAUSS versions 8 and 11 respectively.

The table three above presents the estimated results of the Granger non-causality, modified *WALD* test statistics and the leverage bootstrapped simulated critical values based on the underlying empirical distribution of the data employed in this study. The result of non-Granger causality in column two of the above table indicates the rejection of the null hypothesis of Granger non-causality running from inflation to exchange rate in all countries with only one feedback causation from Nigeria. It means that any policy adjustment on inflation can affect exchange rate stability in all the countries, whereas, policy alteration on exchange rate can influence inflation only in the case of Nigeria. However, the non-Granger causality was developed on asymptotic distribution theory which is seen to results to a spurious conclusion. The null hypothesis at level estimation suffer from non-standard asymptotic distribution, whereas, the integrated Granger causality suffer from independence of nuisance parameter estimates (Sim, Stock & Watson, 1990 and Toda & Philips, 1993).

The modified *WALD* test result shows entirely different results from the Granger non-causality except for Malaysia where the direction of the causation remains the same. Considering Toda and Yamamoto (1995) *MWALD* test result, there exist one way causation from inflation to real exchange rate in Malaysia and Nigeria. While in the case of Philippines and South Africa, feedback causality is obtained. Nonetheless, Sim, Stock and Watson (1990) against the application of asymptotic distribution theory on a VAR model to test for causality. It leads to size distortion under small sample size. More so, the violation of the normality assumption and the presence of ARCH effect in some countries under study make both Granger and *MWALD* tests inefficient (Hacker & Hatemi-J, 2006).

The leverage bootstrapped simulated critical values also indicates the rejection of the null hypothesis of non-Granger causality from inflation to real exchange rate at 1% and 5% for Malaysia and

Nigeria respectively without feedback causality from real exchange rate in these countries. This finding is in line with Cassel (1921) who argue that causality runs from inflation to real exchange rate. The result reveals that policy intervention on inflation can stabilize the Malaysian and Nigerian real exchange rates but not vice versa. Furthermore, bidirectional causation exists in Philippines and South Africa meaning that any policy intervention formulated on one variable can stabilize the other variable. In this case, when policy is formulated on inflation the resultant implication is that it can adjust the exchange rate and vice versa. This result is in line with previous findings such as Arabi (2012) & Madesha, Chidoko and Zivanomoyo (2013).

6. Conclusions

The study employs the more rigorous and robust methods of investigating causality proposed by Toda and Yamamoto (1995) and leverage bootstrap distribution theory. The results are compared between the traditional Granger causality, the modified Wald statistics and the leverage bootstrapped simulated critical values. The contradictory findings obtained from the estimates prove the existence of size distortion and nuisance parameter estimates when the former method is applied. This is usually the case when autoregressive conditional heteroscedasticity exist; integrated series are used in the investigation, and normality of the empirical data is disregarded. The empirical finding based on the simulated critical values indicates the ability of the inflation rate to Granger cause the real exchange rates of Malaysia and Nigeria without being influence by the letter. The result also shows the existence of two ways causation in Philippines and South Africa.

The policy implication of this finding is that the policy makers can manipulate the rate of inflation to stabilize real exchange rate fluctuations in all countries under study, but can only regulate inflation through exchange rate in the case of Philippines and South Africa. This means that monitoring the rate of inflation in these countries can regulate the level of instability in the real exchange rates, whereas, monitoring real exchange rates can only stabilize inflationary pressure in Philippines and South Africa without such conclusion for Malaysia and Nigeria.

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